Interest Parity at Short and Long Horizons *

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Abstract

The unbiasedness hypothesis -- the joint hypothesis of uncovered interest parity (UIP) and rational expectations -- has been almost universally rejected in studies of exchange rate movements. In contrast to previous studies, which have used short-horizon data, we test this hypothesis using interest rates on longer-maturity bonds for the G-7 countries. The results of these long-horizon regressions are much more positive — all of the coefficients on interest differentials are of the correct sign, and almost all are closer to the predicted value of unity than to zero. We first appeal to an econometric interpretation of the results, which focuses on the presence of simultaneity in a cointegration framework. We then use a macroeconomic model to provide an economic explanation for the differences between the short- and long-horizon results. Regressions run on model-generated data replicate the important regularities in the actual data, including the sharp differences between short- and long-horizon parameters. In the short run, the failure of the unbiasedness proposition results from the interaction of stochastic exchange market shocks with endogenous monetary policy reactions. In the long run, in contrast, exchange rate movements are driven by the "fundamentals," leading to a relationship between interest rates and exchange rates that is more consistent with UIP.

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I. INTRODUCTION

Few propositions are more widely accepted in international economics than that the "unbiasedness hypothesis" -- the joint hypothesis of uncovered interest parity (UIP) and rational expectations -- is at best a poor, and often perverse, predictor of future exchange rate movements. In a survey of 75 published estimates, Froot (1990) reports few cases where the sign of the coefficient on interest rate differentials in exchange rate prediction equations is consistent with the unbiasedness hypothesis, and not a single case where it exceeds the theoretical value of unity. This resounding unanimity on the failure of the predictive power of interest differentials must be virtually unique in the empirical literature in economics.

A notable aspect of almost all published studies, however, is that the unbiasedness hypothesis has been tested using financial instruments with relatively short maturities, generally of 12 months or less. There appear to be (at least) three reasons for this practice. The first is constraints on sample size, given that generalized exchange rate floating began only in the early 1970s. This was particularly problematic in the early 1980s, when the floating-rate period was shorter than the maturity of longer-dated financial instruments. The second is that longer-term, fixed-maturity interest rate data were difficult to obtain. The third reason is that some pioneering studies were also concerned with testing the hypothesis of *covered* interest parity, which required observations on forward exchange rates of the same maturity as the associated financial asset. In the event, relatively thick forward exchange markets only exist to a maximum horizon of 12 months.

Fortunately, the length of the floating-rate period is now much longer than when the initial studies were performed, and the availability of data on yields of comparable longer-dated instruments across countries has increased. Accordingly, this paper tests the unbiasedness hypothesis using instruments of considerably longer maturity than those employed in past studies. Our results for the exchange rates of the major industrial countries differ strikingly from those obtained using shorter horizons. For instruments with maturities ranging from 5 to 10 years, *all* of the coefficients on interest rate differentials in the unbiasedness regressions are of the correct sign. Furthermore, *almost all* of the coefficients on interest rates are closer to the predicted value of unity than to the zero coefficient implied by the random walk hypothesis. Finally, as the "quality" of the bond yield data in terms of their consistency with the requirements underlying UIP increases, the estimated parameters typically become closer to those implied by the unbiasedness hypothesis.

To explain the apparently anomalous differences in tests of the unbiasedness hypothesis using short- and long-horizon data, we develop a small macroeconomic model that extends the framework of McCallum (1994a). In particular, a more general monetary reaction function is incorporated that causes interest rates to respond to innovations in output and inflation, as opposed to the exchange-rate targeting framework used by McCallum. Stochastic simulations of the model based are used to generate a synthetic database for replicating unbiasedness tests. Standard regressions using these synthetic data yield negative coefficients on short-term interest rates of roughly the same magnitude as found in most short-horizon studies, thus the model can explain the "forward discount bias" found in such studies. Long-horizon regressions, in contrast, yield coefficients close to unity, consistent with our estimation results using actual data. The long-horizon results differ sharply from the short-horizon results because the model's "fundamentals" play a more important role in tying down exchange rate movements over longer horizons. More generally, the data generated by the simulations for the endogenous variables mimic remarkably closely the key properties of actual data for the G-7 countries.

The paper is structured as follows. Section II reviews the unbiasedness hypothesis, summarizes the existing evidence over short horizons, and provides updated results from 1980 through early 1998. Section III presents estimates of the unbiasedness hypothesis using data on government bond yields for the G-7 countries. Section IV provides an econometric rationalization for the results that are obtained, while Section V develops a model that is consistent with the key features of the observed data and the econometric explanation. Section VI provides concluding remarks.

II. A Review of the UIP Hypothesis and Short-Horizon Evidence

It is convenient to introduce notation and concepts by starting with the covered interest parity (CIP) condition, which follows from the assumption of arbitrage between spot and forward foreign exchange markets. If the conditions for risk-free arbitrage exist, the ratio of the forward to the spot exchange rate will equal the interest differential between assets with otherwise similar characteristics measured in local currencies.[1] Algebraically, CIP can be expressed as:

$$F_{t,t+k} \;/\; S_t \;=\; I_{t,k} \;/\; I^*_{t,k} \;, \tag{1}$$

where S_t is the price of foreign currency in units of domestic currency at time t, $F_{t,t+k}$ is the forward value of S for a contract expiring k periods in the future, $I_{t,k}$ is one plus the k-period yield on the domestic instrument, and $I^*_{t,k}$ is the corresponding yield on the foreign instrument. Taking logarithms of both sides (indicated by lower-case letters), equation (1) becomes:

$$f_{t,t+k} - s_t \;=\; (i_{t,k} - i^*_{t,k}) \;. \tag{2}$$

Equation (2) is a risk-free arbitrage condition that holds regardless of investor preferences. To the extent that investors are risk averse, however, the forward rate can differ from the expected future spot rate by a premium that compensates for the perceived riskiness of holding domestic versus foreign assets. We define the risk premium, η, accordingly:

$$f_{t,t+k} \;=\; s^e_{t,t+k} + \eta_{t,t+k} \;. \tag{3}$$

Substituting equation (3) into (2) then allows the expected change in the exchange rate from period t to period $t+k$ be expressed as a function of the interest differential and the risk premium:

$$\Delta s^e_{t,t+k} \;=\; (i_{t,k} - i^*_{t,k}) - \eta_{t,t+k} \;, \tag{4}$$

Narrowly defined, UIP refers to the proposition embodied in equation (4) when the risk premium is zero—consistent, for instance, with the assumption of risk-neutral investors. In this case, the expected exchange rate change equals the current interest differential. Equation (4) is

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[1] These conditions include identical default risk and tax treatment, the absence of restrictions on foreign ownership, and negligible transactions costs.

not directly testable, however, in the absence of observations on market expectations of future exchange rate movements.[2] To operationalize the concept, UIP is generally tested jointly with the assumption of rational expectations in exchange markets. In this case, future realizations of s_{t+k} will equal the value expected at time t plus a white-noise error term $\xi_{t,t+k}$ that is uncorrelated with all information known at t, including the interest differential and the spot exchange rate:

$$s_{t+k} \;=\; s^{re}_{t,t+k} \;+\; \xi_{t,t+k} \;, \tag{5}$$

where $s^{re}_{t,t+k}$ is the rational expectation of the exchange rate at time $t+k$ formed in time t. Substituting equation (5) into (4) gives the following relationship:[3]

$$\Delta s_{t,t+k} \;=\; (i_{t,k} \;-\; i^{*}_{t,k}) \;-\; \eta_{t,t+k} \;+\; \xi_{t,t+k} \;, \tag{6}$$

where the left-hand side of equation (6) is the realized change in the exchange rate from t to $t+k$.

It is natural, then, to test the composite "unbiasedness" hypothesis of UIP and rational expectations via the regression equation:

$$\Delta s_{t,t+k} \;=\; \alpha \;+\; \beta\,(i \;-\; i^{*})_{t,k} \;+\; \epsilon_{t,t+k} \;. \tag{7}$$

Under the assumption that the composite error term $\epsilon_{t,t+k}$ is orthogonal to the interest differential, the estimated slope parameter in equation (7) should be unity. In addition, no other regressors known at time t should have explanatory power, as all available information should be captured in

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[2] Indirect tests of UIP have been performed using surveys of published forecasts of exchange rates. Chinn and Frankel (1994a,b) find mostly positive correlations between the forward discount and the expected depreciation, which is consistent with UIP.

[3] This condition is also sometimes referred to as the "risk-neutral efficient markets hypothesis." In the absence of risk neutrality, market efficiency does not require that the forward exchange rate equals its expected future level. Tests of this more general version of market efficiency are not possible, however, in the absence of direct measures of risk premia in exchange markets.

the rational expectation of $\Delta s_{t,t+k}$ as reflected in the period-t interest differential. Regarding the constant term, non-zero values may still be consistent with UIP. Jensen's inequality, for instance, implies that the expectation of a ratio is not the same as the ratio of the expectations (see Meese,1989).[4] Alternatively, relaxing the assumption of risk-neutral investors, the constant term may reflect a constant risk premium demanded by investors on foreign versus domestic assets. Default risk could play a similar role, although the latter possibility is less familiar because tests of UIP (as well as CIP) generally use returns on assets issued in offshore markets by borrowers with comparable credit ratings. In contrast, the long-term government bonds used for estimation in Section III may not share the same default attributes, so that a pure default risk premium might exist.

As noted above, estimates of equation (7) using values for k that range up to one year resoundingly reject the unbiasedness restriction on the slope parameter. The survey by Froot (1990), for instance, finds an average estimate for β of -0.88.[5] Thus, the common perception that the failure of UIP indicates that short-run exchange rate movements are best characterized as a random walk is not strictly true: over short horizons, most studies find that exchange rates move *inversely* with interest differentials.

To illustrate the dismal performance of short-horizon interest rates as predictors for movements in the exchange rates of the G-7 countries, Table 1 presents estimates of equation (7) for the period from the first quarter of 1980 to the first quarter of 1998. The exchange rates of the other six countries were expressed in terms of U.S. dollars, and the 3-, 6-, and 12-month movements in exchange rates were regressed against differentials in eurocurrency yields of the

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[4] As noted in Engel (1996), however, a constant term due to Jensen's inequality is likely to be small in practice.

[5] Other recent surveys that report similar results include MacDonald and Taylor (1992), Isard (1995) and Lewis (1995). A qualified exception is the study by Flood and Rose (1996), which finds that the coefficient on the interest differential is closer to its UIP value during periods when exchange rate realignments within the ERM were expected (and observed).

corresponding maturity.[6] Estimation using the 6- and 12-month horizon data at a quarterly frequency led to overlapping observations, inducing (under the rational expectations null hypothesis) moving average (MA) terms in the residuals. Following Hansen and Hodrick (1980), we used the Generalized Methods of Moments (GMM) estimator of Hansen (1992) to correct the standard errors of the parameter estimates for moving average serial correlation of order k-1 (i.e., MA(1) in the case of 6-month data and MA(3) in the case of 12-month data).

The results confirm the failure of UIP over short horizons, similar to other studies. At each horizon, four of the six estimated coefficients have the “wrong” sign relative to the unbiasedness hypothesis. The average coefficient is around -0.8, similar to the value in the survey by Froot (1990). Panel estimation with slope coefficients constrained to be identical across countries yields estimates ranging from about -0.6 at the 6-month horizon to -0.3 at the 12-month horizon. In most cases it is possible to reject the hypothesis that β equals unity; in cases where UIP cannot be rejected, the standard errors of the estimated parameters are sufficiently large that it would be difficult to reject almost any plausible hypothesis.Only for the lira is it possible to reject the random-walk model while not also rejecting UIP.[7] All of the adjusted R^2 statistics (not reported) are very low, and occasionally negative.

In Figure 1, recursive coefficient estimates for β are displayed for all six exchange rates. It is clear that some point estimates wander considerably (Canadian dollar, Deutschemark), while others do not. In particular, the point estimate for the lira appears consistently positive. Another interesting result is obtained if one breaks the 1980Q1-1998Q1 sample at 1979Q4/1980Q1; the UK point estimate for the latter subsample is positive, while it is negative in the early subsample.

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[6] Yields and exchange rates were both constructed as the average of bid and offer rates on the last trading day of each quarter. Exchange rate movements and interest differentials are expressed at annual rates.

[7] Although one cannot formally test the null of a zero coefficient, since the standard errors are constructed under the null hypothesis that β=1.

III. LONG-HORIZON ESTIMATES

As noted in the introduction, short-horizon tests of the unbiasedness hypothesis have been facilitated by the availability of interest rate series that correspond closely to the requirements for covered interest parity. Data of comparable quality for longer-horizon instruments generally are much less readily available. In particular, it is difficult to obtain longer-term rates in offshore markets on thickly-traded instruments of a known fixed maturity. For the purposes of this study, then, we have used data that are inherently somewhat less pure from the point of view of the UIP hypothesis. Specifically, we have used domestic yields on government bonds of maturities that correspond roughly, but not necessarily exactly, to assumed constant maturities. Moreover these on-shore instruments may be subject to differences in tax regime, capital controls, etc. Nonetheless, based on the findings by Popper (1993) that covered interest differentials at long maturities are not appreciably greater than those for short (up to one year) maturities, we do not expect that rejections of long-horizon UIP will be driven by deviations from covered interest parity.

Even if these data tend to exhibit more "noise" than those used for short-horizon tests of UIP, for conventional errors-in-variables reasons we would expect the coefficient on the interest differential in these long-horizon regressions to be biased *toward* zero, and away from its hypothesized value of unity. Hence, the results we obtain should be conservative in nature.

The first data set we employ to test long-horizon unbiasedness consists of the benchmark government bond yields used by Edison and Pauls (1993), provided by Hali Edison. These are end-of-month yields on outstanding government bonds for the G-7 countries, generally of 10-year maturity at the date of issuance. The 10-year change in the exchange rate versus the dollar for the other six currencies is then regressed on the 10-year lagged differential in the associated bond yield.[8] Given that floating rates were generally introduced in 1973, after allowing for the 10-year

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[8] The serial correlation problem becomes a potentially serious issue as the number of overlapping observations increases rapidly with the instrument maturity. One way to overcome the problem is to use only non-overlapping data; however, this procedure amounts to throwing away information. Boudoukh and Richardson (1994) argue that, depending upon the degree of serial correlation of the regressor and the extent of the overlap, using (continued...)

lag on the interest differential, the available estimation period consisted of 1983Q1–1998Q1 (given limitations on the availability of bond yield data for Italy, the sample period for the lira begins in1985Q1).

The results of these regressions are reported in the first panel of Table 3. They represent a surprising and stark contrast to the short-horizon results reported in Section II. In all cases, the estimated slope coefficient is positive, with four of the six values lying closer to unity than to zero. For the Canadian dollar, the point estimate (1.104) is very close to unity, while the deutschemark and the franc also evidence high coefficients. The yen, pound and lira are the three cases in which UIP is statistically rejected. But for all currencies except the lira, the hypothesis that β equals zero can also be strongly rejected. The adjusted R^2 statistics are also typically higher than in a typical short-horizon regressions, with the proportion of the explained variance in the deutschemark and the pound approaching one half.

Since there are relatively few independent observations in the single-currency regressions, additional power can be obtained by pooling the data and constraining the slope coefficient to be the same across currencies.[9] The resulting point estimate is reported under the entry “constrained panel” at the bottom of Table 2.a. Its value of 0.645 is well below unity; on the other hand, it is closer to unity than to zero, a substantial difference from the panel estimates obtained for horizons up to one year reported in Table 1.

For Japan, Germany, the U.K., and the U.S., it was also possible to obtain synthetic “constant maturity” 10-year yields from interpolations of the yield curve of outstanding government securities. The regressions using measures of long-horizon interest differentials based on these data are reported in Table 2.b. For each of the three associated exchange rates, the estimated slope parameter is closer to unity than in the regressions using benchmark yields

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[8](...continued)
overlapping data is equivalent to using between 3 to 4.5 times the number of observations available otherwise.

[9] These are fixed effects regressions which allow for a different constant across currencies. The standard errors are constructed to allow for cross-currency correlations, as well as serial correlation due to overlapping horizons. See Frankel and Froot (1987) for details.

(although the difference is slight for the deutschemark). Moreover, the panel point estimate of 0.708 is also closer to the posited value. The improvement in the results, although modest, suggests that part of the reason why unbiasedness is still rejected when using benchmark yields relates to discrepancies between the assumed and actual maturities of the outstanding securities. In other words, improvements in the quality of the data appear to systematically shift the results toward supporting the UIP hypothesis.

Similar constant-maturity 5-year yields were obtained for Germany, the U.K., Canada, and the U.S. Regressions of 5-year changes in exchange rates on the interest differentials implied by these data are reported in Table 2.c. The results are even more favorable for the UIP hypothesis: for both the deutschemark and the pound, the slope coefficients are economically and statistically indistinguishable from the implied value of unity, while the null of zero under the random walk hypothesis is strongly rejected. In the case of the Canadian dollar, the point estimate is 1.34, but the 50 percent confidence bound for the point estimate easily encompasses the value of unity. Similarly, the panel estimate is essentially equal to the theoretically implied value.

The only other studies that we are aware of that test the unbiasedness hypothesis over horizons of longer than 12 months are Flood and Taylor (1997), and Alexius (1999). Flood and Taylor regress 3-year changes in exchange rates on annual average data on medium-term government bonds from the IMF's *International Financial Statistics* (*IFS*). The data over the 1973–92 period are then pooled for a sample of 21 countries. They find a coefficient on the interest differential of 0.596 with a standard error of 0.195. Thus the hypotheses that β equals either zero or unity can both be rejected. These results are broadly in line with our 10-year results, although our 5-year results using constant maturity data are more supportive of the unbiasedness hypothesis. This difference may reflect the fact that our end-period, constant-maturity data are more closely aligned with the requirements underlying the UIP hypothesis, although differences in country coverage and sample periods may also play a role.

In the study by Alexius, 14 long term bond rates (of uncertain maturities) for the 1957Q1-199Q4 period are drawn from *IFS*. She attempts to control for the measurement error arising from uncertain maturities, and the role of coupon payments.[10] Her study also finds substantial evidence in favor of the unbiasedness hypothesis at long horizons. For the deutschemark, the OLS point estimate for the duration- and coupon-adjusted series is 0.820, which is remarkably close to our estimate of 0.836 for the 10 year constant maturity yields. On the other hand, her estimates for the yen and the pound (0.209 and 0.278, respectively) are somewhat lower than the estimates we report in Table 2.b of 0.564 and 0.719. Some of this difference may be due to the longer sample she uses, which encompasses a period of substantial capital controls.

In any event, it is reassuring that the Flood and Taylor and Alexius results are similar to those obtained in our regressions, suggesting that the difference between short- and long-horizon tests of UIP may be robust across countries, sample periods and estimation procedures.

IV. Explaining the Results Econometrically

The rather strikingly different results obtained at different horizons should be placed in the context of recent findings that when the unbiasedness proposition is couched in terms of cointegrating relationships, one finds that it is much more difficult to reject the null hypothesis of unbiasedness (e.g., Evans and Lewis, 1995). Here, we are not so much concerned with the specific finding regarding cointegration with the posited values, but rather the econometric implications of estimating equation (7). If the expected spot and forward rate are cointegrated, then according to the Engle-Granger Representation Theorem, one can re-write the cointegrated system as:

$$\begin{aligned}\Delta s_t &= \gamma_{10} + \Phi_1[s_{t-1} - \delta_1 f_{t-2} - \delta_0] + \sum_{i=1}^{j} \gamma_{1i}\Delta s_{t-i-1} + \sum_{i=1}^{j} \zeta_{1i}\Delta f_{t-i-2} + \epsilon_{1t} \\ \Delta f_{t-1} &= \gamma_{20} + \Phi_2[s_{t-1} - \delta_1 f_{t-2} - \delta_0] + \sum_{i=1}^{j} \gamma_{2i}\Delta s_{t-i-1} + \sum_{i=1}^{j} \zeta_{2i}\Delta f_{t-i-2} + \epsilon_{2t}\end{aligned} \qquad \textbf{(8)}$$

where the forward rate horizon has been set to one (k = 1) for simplicity. As pointed out by Phillips (1991), single-equation estimation of (8.a) is plagued by asymptotic bias as long as the forward rate is not weakly exogenous. To see how this relates to the conventional depreciation-

[10] The *IFS* data are somewhat problematic in that the definitions of the long term bonds is not homogeneous across countries, and change at certain times.

forward discount regression in (7), consider that the forward rate must be cointegrated with the contemporaneous spot rate if it is cointegrated with the future spot rate. Following Moore (1994), rewrite (8.a), assuming weak exogeneity of f ($\Phi_2 = 0$)

$$\begin{aligned} \Delta s_t &= b_0 \Delta f_t + \Phi_1 [s_{t-1} - \delta_1 f_{t-1} - \delta_0] \\ &\quad + \sum_{i=1}^{j-1} b_i \Delta s_{t-i} + \sum_{i=1}^{j-1} c_i \Delta f_{t-i,1} + u_t \end{aligned} \tag{9}$$

where b_i and c_i are functions of the variances and covariances of ϵ_1 and ϵ_2, and u is a function of ϵ_1 and ϵ_2, and their variances and covariances. In particular, $b_0 = \sigma_{12}/\sigma_{22}$, which equals zero only when the correlation between the ϵ's is zero. Imposing the restrictions $\delta_0 = 0$ and $\delta_1 = 1$ in the cointegrating vector,[11] equation (9) can be rewritten as:

$$\begin{aligned} \Delta s_t &= -\Phi_1 \delta_0 - \Phi_1 [f_{t-1} - s_{t-1}] + b_0 \Delta f_t + \Phi_1 (1 - \delta_1) f_{t-1} \\ &\quad + \sum_{i=1}^{j-1} b_i \Delta s_{t-i} + \sum_{i=1}^{j-1} c_i \Delta f_{t-i} + u_t \end{aligned} \tag{10}$$

Notice that equation (10) degenerates to equation (7) if and only if $b_0 = 0$, $b_i = c_i = 0$ for all i, *and* $\delta_1 = 1$ (Moore, 1994; Villanueva, 1998).

To examine whether the standard assumption of weak exogeneity of the forward rate is justified at either the short or long horizons, we generate implicit forward rates using the exact relationship in equation (1), for both the 3 month and 5 year horizons. We then test for cointegration between the forward rate and the future spot rate[12] using the Johansen (1988)

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[11] Although it is common to impose the zero constant in a joint test of the coefficients, Brenner and Kroner (1995) have shown that when the (log) spot and forward rates follow continuous time random walk processes, then there will be a constant in the cointegrating relationship even under the risk neutral assumption. This constant will be a function of the drift term and the variance of innovations in the underlying process driving the spot rate.

[12] In principle, either specification is valid asymptotically. Zivot (1997) argues for testing the cointegrating vector involving the contemporaneous forward and spot rate, while Villanueva (1998) reports results demonstrating that lagged forecast errors yield more unambiguous results.

maximum likelihood procedure. The results are reported in Table 3; in Panel 3.a are the cointegration results for the 3 month forward rates and the future spot rates, and in Panel 3.b are the corresponding results for the 5 year implicit forward rates.

The first column displays the likelihood ratio for the Maximal Eigenvalue statistic. The 5% critical value for rejecting the null hypothesis of no cointegrating vectors in favor of the alternative of one is 15.41. The lira, yen and pound clearly evidence cointegration, while the franc test statistic is borderline significant. In the second and third columns, δ_0 is the constant in the cointegrating vector, and δ_1 is the slope coefficient. Under the maintained hypothesis of long run unbiasedness, $\delta_1 = 1$. The results indicate that for the yen, pound, franc and lira, the spot rate responds to the lagged error correction term. For the lira, yen and pound, the forward rate also responds. Among the cases for which cointegration is detected, only the franc exhibits weak exogeneity of the forward rate. Interestingly, this is a case for which the 3 month forward rate is *not* an unbiased predictor of the future spot rate.

For the 5 year implicit forwards and the corresponding future spot rates, borderline cointegration is detected for the pound, and somewhat less evidence is detected for the deutschemark. The Canadian sample is much shorter than that for the other currencies, so we are not able to find formal evidence of cointegration. Interestingly, for the pound, one obtains the result that at horizons of 5 years, the spot rate responds to the lagged cointegrating vector Φ_1 with high statistical significance, while the forward rate does not. That is, long term interest rate differentials are weakly exogenous in this system.

Unfortunately, the cointegration evidence for the deutschemark is weaker. If one uses the more powerful Horvath-Watson (1995) test, which imposes the $\delta_1 = 1$ restriction, one finds that test statistic of 4.52 is less than the 10% critical value of 8.30 for the case with a nonzero mean in the variables (although it is only slightly less than the corresponding critical value of 4.73 for the less relevant zero-mean case). If one were willing to impose the prior of cointegration (see Kremers, Ericsson and Dolado, 1992), then the t-statistic on Φ_1 is 1.87, while that on Φ_2 is only 1.00. The data thus seem to suggest that the 5 year deutschemark forward rate is weakly exogenous.

For the Canadian dollar, no evidence of cointegration can be detected. The results in the Panel 3.b are for the Horvath-Watson regressions where the null of long run unbiasedness is imposed. The spot rate appears to be more responsive to the forward forecast error than the forward rate. However, these latter results are merely suggestive because we are not able to detect cointegration in the (admittedly short) sample we have.[13]

For two of the three currencies for which we have data, it appears that the forward rate is weakly exogenous at long horizons, while at short horizons the spot rate is more likely to be weakly exogenous. From a purely statistical standpoint, this explains some of the differences in the results obtained at short and long horizons. The challenge then remains to determine an economic reason why this should be the case.

IV. A Macroeconomic Interpretation

Here, we propose a solution to the unbiasedness puzzle based on the properties of a small macroeconomic model that incorporates feedback mechanisms between exchange rates, inflation, output, and interest rates. In particular, the model generates simulated data that are fully consistent with the stylized facts: that regressions using short-horizon data yield negative slope coefficients and explain little if any of the variance in exchange rates, while long-horizon regressions yield coefficients close to unity and explain a much higher proportion of exchange rate movements.

The model is in the spirit of the framework outlined in McCallum (1994a), but allows for a richer interaction between interest rates and exchange rates. Stochastic simulations of the model are performed to generate a synthetic database, which is then used to replicate standard short- and long-horizon tests of UIP. The regressions using the synthetic data are similar to those obtained using actual data for the G-7 countries, with a pronounced difference between the short- and long-horizon parameter estimates. In the short run, shocks in exchange markets lead to monetary

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[13] The Canadian constant 5 year maturity yield series begins only in1980Q4. Another interesting aspect of the Canadian dollar is the large Canada-US interest differential which appeared in 1990 with the collapse of the Meech Lake accords, and disappears in 1997 (see Clinton, 1998).

policy responses that result in a negative correlation between exchange rates and interest rates, contrary to the unbiasedness hypothesis under UIP. Over the longer term, in contrast, exchange rates and interest rates are determined by the macroeconomic “fundamentals” of the model, and thus behave in manner more consistent with the conventional UIP relationship.

McCallum’s framework is based on a two-equation system consisting of an uncovered interest parity relationship augmented by a monetary reaction function that causes interest rates to move in response to exchange rate changes:

$$\Delta s^{e}_{t,t+1} = (i_t - i_t^*) - \eta_t$$

$$(i_t - i_t) = \lambda \, \Delta s_t + \sigma (i_{t-1} - i_{t-1}^*) + \omega_t \ ,$$

where i_t -i_t^* represents the interest differential, η_t is a stochastic shock to the uncovered interest parity condition, and ω_t is an interest rate shock. McCallum is agnostic about the nature of the factors that underly η. We follow the same convention, simply calling it for the time being an “exchange market” shock. McCallum solves this model to show that the parameter on the interest rate in the reduced-form expression for the change in the next-period exchange rate is $-\sigma/\lambda$, which will be negative given conventional parameter values.[14]

The applicability of McCallum’s interest rate reaction function has been criticized by Mark and Wu (1996), who find a value of λ that is small and insignificant for Germany, Japan, and the U.K. More generally, his reaction function does not incorporate variables that are usually believed to be of concern to policymakers, such as inflation and output. In this sense, McCallum’s model does not provide a complete characterization of macroeconomic interactions, but serves the narrower purpose of illustrating how a negative correlation between interest rates and exchange rate movements might be generated in a consistent framework.

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[14] He also allows for first-order autocorrelation in η. In this case, the parameter on the interest rate becomes $(\rho-\sigma)/\lambda$, which McCallum argues will also be negative for plausible parameter values.

To generalize McCallum's model, and allow a richer characterization of the feedback process between interest rates and exchange rates, we extend it by including equations for output and inflation. The monetary reaction function is then specified so that interest rates adjust in response to movements in output and inflation, using the rule proposed by Taylor (1993). To the extent that output and inflation are affected by the exchange rate, interest rates will still respond to innovations in the disturbance in the UIP relationship, but through a less direct channel than originally posited by McCallum. The model is described in Table 4, where the variables are interpreted as being measured relative to those in the partner country against which the exchange rate is defined—in this case, the United States. The periodicity is assumed to be annual, and all variables are expressed at annual rates.

The inflation equation is an expectations-augmented Phillips curve: current period inflation adjusts in response to past inflation, expected future inflation, the current output gap, and the current change in the real exchange rate. The theoretical justification for this type of equation is discussed in Chadha, Masson, Meredith (1993). Parameter values have been chosen to be broadly consistent with the empirical evidence using panel data for the G-7 countries. The output equation is a standard open-economy IS curve, where output responds to the real exchange rate, the expected long-term real interest rate, and the lagged output gap. The parameters have been chosen such that a 10 percent appreciation in the real exchange rate reduces output by 1 percent in the first year, and by 2 percent in the long run; a 1 percentage point rise in the real interest rate lowers output by ½ percent in the first year and 1 percent in the long run.[15] The long-term interest rate is determined as the average of the current short-term interest rate and its expected value over the four subsequent periods—thus, the long-term rate can be thought of as a five-year bond yield that is determined by the expectations theory of the term structure. Expected long-term inflation is defined similarly in constructing the real long-term interest rate.

Stochastic elements are introduced via three processes, all of which are assumed to be white noise: exchange market shocks (η_t), inflation shocks (ν_t), and output shocks (ϵ_t). We

[15] These responses are broadly consistent with the average values across the G-7 countries embodied in MULTIMOD, the IMF's macroeconomic simulation model (Masson, Symansky, and Meredith (1990)).

characterize the solution using numerical simulations based on the stacked-time algorithm for solving forward-looking models described in Armstrong, Black, Laxton, and Rose (1998).[16] An important feature of the solution path is that expectations are consistent with the model's prediction for future values of the endogenous variables, based on available information about the stochastic processes. As the innovation terms η_t, ν_t, and ϵ_t are assumed to be independent and uncorrelated, the information set consists of the contemporaneous innovations as well as the lagged values of the endogenous variables. At any point in time, the conditional expectation of the future values of the innovations is zero given the assumption that they are white noise. In this sense, expectations are fully rational given the model structure. Nevertheless, agents lack perfect foresight, because they cannot anticipate the sequence of future innovations that determine the realizations of the endogenous variables. As the innovations are white noise, so are the associated expectational errors.

The only other information needed to perform the simulations is the relative variance of the three stochastic processes. These were chosen to yield simulated variances of exchange rates, inflation, and output that are consistent with the stylized facts for the G-7 countries. Specifically, the standard deviation of the year-to-year movement in the exchange rate, averaged across the G-7 countries (excluding the U.S., the numeraire currency) is about 12.0 percent for the 1975-97 period. The standard deviations in the year-to-year movements in inflation and output (relative to the US) are much lower, at about 2.0 percent and 1.9 percent respectively. Experimental simulations indicated that these values were broadly consistent with a standard deviation for the exchange market innovation of 9.7 percentage points, for the inflation innovation of 1.3 percentage points, and for the output innovation of 1.1 percent.

To illustrate the model's properties, Figure 1 shows impulse responses for standardized innovations in each disturbance. In the face of a temporary exchange market shock, the exchange rate depreciates by 9 percent in the first period. This raises inflation by almost 1 percent, and

[16] The performance of this algorithm is compared with that of other forward-looking solution methods in Juillard, Laxton, McAdam, and Pioro (1998). The simulations were performed using Portable Troll version 1.031. Data and programs are available on request from the authors.

output by ¾ percent. Under the Taylor Rule, these movements in inflation and output cause the short-term interest rate to rise by slightly over 1½ percentage points. In the second period the shock dissipates and the exchange rate *appreciates* by about 8 percent, reversing the initial increase in inflation and the short-term interest rate, while output declines toward its baseline level. The exchange rate appreciation in the second period occurs in spite of a higher lagged short-term interest rate, implying the opposite response to that predicted by UIP. This reflects the rise in the lagged exchange market shock, which generates a perverse short-run correlation between the lagged interest rate and the next-period change in the exchange rate. From a low-frequency perspective, though, the effects of the exchange market shock show little persistence. This is reflected in the muted response of the long-term interest rate (defined here as the 5-year bond yield), which increases by only ¼ percentage point in the first period before returning close to baseline in the second.

An inflation shock causes the short-term interest rate to rise by roughly the same amount in the first period as an exchange market shock. The exchange rate initially appreciates in response to higher interest rates, followed by depreciation in subsequent periods, as implied by the well-known "overshooting" model of Dornbush (1976).[17] In all periods after the first period (when the shock hits), the relationship between the change in the exchange rate and the lagged interest rate is consistent with UIP, in contrast to the situation with an exchange market shock. The long-term interest rate also initially rises by much more, indicating that the inflation shock has greater persistence in its effects on short-term interest rates. This difference is important, because it implies a greater covariance between the long-term interest rate and the future change in the exchange rate under an inflation shock than under an exchange market shock. This, in turn, puts greater weight on comovements in interest rates and exchange rates that are UIP-consistent at longer horizons.

Similarly, an output shock causes short- and long-term interest rates to rise on impact, while the exchange rate initially appreciates followed by subsequent depreciation. Although the

[17] The long-run depreciation of the nominal exchange rate under an inflation shock reflects an increase in the domestic price level, which is not tied down under the Taylor Rule. The *real* exchange rate returns to its initial level in the face of a temporary inflation shock.

changes in interest rates are not as large as under an inflation shock, the results are qualitatively similar—long-term interest rates rise by much more than with an exchange market shock, which again results in more weight being placed on UIP-consistent movements in the data at longer horizons.

To confirm the intuition provided by the impulse response functions, stochastic simulations were performed on the model. Each simulation was performed over a 140-year horizon, with the first 30 and last 30 years being discarded to avoid contamination from beginning- and end-point considerations. This yielded a “sample” of 80 observations for each simulation. This process was repeated 50 times to generate a hypothetical population of 50 such samples. For each sample, standard UIP regressions were run using horizons varying from 1 to 10 years. The results of the 1-year and 5-year regressions for a representative population of 50 simulations are shown in Table 5.

The most prominent feature of the results is the difference in the slope parameters between the regressions at the 1-year horizon versus those at longer horizons. For the 1-year regressions, the average slope parameter of -0.50 is the same order of magnitude as those obtained in Section II using data for the G-7 countries. Given the average standard error of 0.42, it would easily be possible to reject the hypothesis that β equals unity with a high level of confidence in the typical sample. In both the 5-year and 10-year regressions, the average estimated β is 0.82, with a standard error of only 0.18. Thus, one could reject the hypothesis that β equals zero at conventional confidence levels, but not generally reject the hypothesis of unity. These results are generally consistent with the pooled regressions using long-horizon data reported in Section III. There are large outliers in some of the samples, however. The short-horizon coefficient ranges from -1.37 to 0.37 across samples, indicating the variability in estimation results that could be obtained using samples even as long as 80 periods. The 5-year results are somewhat more tightly clustered, ranging from 0.53 to 1.32. For both the short- and long-horizon regressions, the average standard error of β found in individual samples is similar to the standard deviation calculated across the 50 samples, suggesting that the calculated standard errors in the regressions are indeed good estimates of the sample variability of the coefficients.

Another interesting comparison between the regressions involves the adjusted R^2 statistics. The average value in the 1-year regressions is only 0.01, indicating a virtual complete lack of explanatory power, similar to the regressions using actual data. For the 5-year regressions, in contrast, the average value rises to 0.21, with some draws as high as 0.46. Again, this is consistent with the stylized facts from the actual long-horizon regressions reported above. Thus, even in the context of a model whose structure is unchanging over time and where agents are assumed to have fully rational expectations, interest differentials do not explain the bulk of the variance in longer-term exchange rate movements. This reflects the influence of future innovations that are inherently unpredictable in affecting the future exchange rate path.

Figure 2 illustrates the pattern of the slope coefficients at alternative horizons for three different populations of simulations. They yield very similar results, with the average slope coefficients ranging from -0.4 to -0.5 at the 1-year horizon, but becoming significantly positive at horizons of 2 years and more. Indeed, the 3-year horizon coefficients are quite close to the values of 0.7–0.8 reached at 5- and 10-year horizons. It is also interesting to note that the coefficients do not asymptote toward unity at these longer horizons, but rather stabilize at levels somewhat below the value implied by the unbiasedness hypothesis. This reflects the diminishing—but nonnegligible—role that exchange market shocks continue to play at longer horizons. The implication is that it may be unrealistic to expect to find coefficients centered on unity in UIP tests at any horizon, even in the absence of measurement errors in the data. These results are consistent with those obtained using actual long-horizon data. Flood and Taylor estimate a coefficient at a 3-year horizon of about 0.6, only slightly below that implied by the synthetic regressions. Our 5-year results are actually somewhat more favorable to UIP than implied by the synthetic regressions, but the coefficients differ by less than one standard deviation of the value estimated in the synthetic regressions. Our 10-year results are only slightly below the synthetic value.

In terms of the consistency of the model-generated volatility of the main variables with actual data, the following tabulation compares the mean results across the simulations with average values across the G-7 countries for 1975-97:

	Actual	Simulated
Standard deviation of:		
Δs_t	12.0	12.5
$\Delta\hat{\pi}_t$	2.0	2.0
$\Delta\hat{y}_t$	1.9	1.4
$\Delta\hat{i}_t$	3.4	3.3
$\Delta\hat{i}^l_t$	1.1	0.9
Correlation of:		
$\hat{i}_t$, $\hat{i}_{t-1}$	0.52	0.50

Note: "^" denotes a variable expressed relative to the US. For the actual data, the sample statistics refer to the 1975-97 period.

The model replicates closely the observed volatility in the actual data, although the simulated standard deviations of changes in output and long-term interest rates are somewhat lower than the averages observed for the G-7 countries. It is interesting to note that changes in short-term interest rates exhibit similar volatility in the simulations as in the observed data, even though no explicit interest rate shock is incorporated in the model. In addition, the correlation of short-term interest rates with their lagged values is very similar in the simulations to that in the actual data, in spite of the fact that an "interest-rate smoothing" term is not included in the reaction function and the model's innovations are serially uncorrelated. This reflects the propagation over time of uncorrelated innovations via the lagged dependent variables in the inflation and output equations.[18]

The somewhat lower simulated variance of the long-term interest rate compared with the actual data may reflect the absence of error terms (i.e. risk premia) in the term structure equations, contrary to empirical evidence (see McCallum (1994b)). To check the sensitivity of the results to this assumption, simulations were preformed with additional stochastic

[18] This contrasts with McCallum's model, which requires the assumption of serially correlated exchange market shocks to generate serial correlation in interest rates, even though the model incorporates an interest-rate smoothing mechanism.

disturbances added to the term structure relationships.[19] The short-horizon estimation results were virtually unaffected by this addition, while the slope parameters in the long-horizon regressions declined modestly—for instance, the average value of β in the 5-year regressions fell from 0.82 to 0.78, while that at the 10-year horizon fell from 0.78 to 0.68. This is not surprising, as the term premium introduces what amounts to an error in the regressor in the long-horizon regressions. For conventional reasons, this source of noise would bias the estimated coefficient toward zero. But the magnitude of the effect is modest in the simulations and the estimated parameter remains similar to those obtained with actual data.

Thus it appears that a small, forward-looking macroeconomic model with a conventional structure is capable of explaining important stylized facts relating to tests of UIP. The failure of UIP over short horizons is consistent with the endogeneity of interest rates in the face of disturbances in exchange markets. Over the longer term, in contrast, the model's "fundamentals" dominate and UIP performs better. This interpretation is consistent with the finding in Mark (1995) and Chinn and Meese (1995) that short-horizon movements in exchange rate are dominated by noise while longer-term movements can be related to economic fundamentals.

Nevertheless, this framework cannot explain another major puzzle—the source of the disturbances in exchange markets required to generate the observed volatility in exchange rates. It is well known that conventional consumption-based asset pricing models are unable to generate risk premia of the magnitude required to explain observed price fluctuations, not only in exchange markets, but in almost all financial markets. More recent analyses based on "first-order" risk aversion, such as Bekaert, Hodrick, and Marshall (1997), also generate risk premia that are far smaller than the shocks observed in the data. Beyond this, it has also proved difficult to relate ex post exchange risk premia to macroeconomic factors. It appears that alternatives to approaches based solely on agents' aversion to consumption risk are needed to explain the stylized facts of predictable excess returns in asset markets.

[19] The standard deviations of the disturbances were calibrated to raise the standard deviation of the year-to-year movements in the long-term (10-year) bond yield to match that in the observed data, i.e. 1.1 percentage points.

V. Conclusions

We find strong evidence for the G-7 countries that the perverse relationship between interest rates and exchange rates is a feature of the short-horizon data that have been used in almost all previous studies. Using longer horizon data, the results of standard test of UIP yield strikingly different results, with slope parameters that are positive and closer to the hypothesized value of unity than to zero. These results confirm the earlier conjectures of Mussa (1979) and Froot (1990) that the unbiasedness proposition may better apply at longer horizons.

The difference in the results is shown to be fully consistent with the properties of a conventional macroeconomic model. In particular, a temporary disturbance to the uncovered interest parity relationship causes the spot exchange rate to depreciate relative to the expected future rate, leading to higher output, inflation, and interest rates. Higher interest rates are then typically associated with an ex post future appreciation of the exchange rate at short horizons, consistent with the forward discount bias typically found in empirical studies. Over longer horizons, the temporary effects of exchange market shocks fade and the model results are dominated by more fundamental dynamics that are consistent with the UIP hypothesis. The model, though, cannot explain why such shocks are as large as needed to explain observed exchange rate volatility.

Regardless of the reasons for the failure of the unbiasedness hypothesis at short horizons, from an unconditional forecasting perspective, the conclusion remains that interest differentials are essentially useless as predictors of short-term movements in exchange rates. Over longer horizons, however, our results suggest that interest differentials may significantly outperform naive alternatives such as the random-walk hypothesis, although it is still likely to explain only a relatively small proportion of the observed variance in exchange rates.

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Table 1. Estimates of β

$$\Delta s_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^*_{t,k}) + \epsilon_{t+k}\ . \qquad \textbf{(7)}$$

	Maturity		
Currency	3 mo.	6 mo.	12 mo.
Deutschemark	-0.646* (1.160)	-0.861** (0.839)	-0.581*** (0.682)
Japanese yen	-3.440*** (0.998)	-3.387*** (0.783)	-2.996*** (0.706)
U.K. pound	-2.036*** (1.211)	-1.963*** (1.176)	-1.268** (1.062)
French franc	0.273 (1.010)	0.269 (0.920)	0.491 (0.889)
Italian lire	1.462 (0.858)	1.847 (0.795)	1.994 (0.647)
Canadian dollar	-0.516*** (0.589)	-0.535*** (0.391)	-0.464** (0.477)
Panel[1]	-0.507*** (0.405)	-0.568*** (0.390)	-0.321*** (0.423)

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample for 3, 6, and 12 month maturity data is 1980Q1-1998Q1. * (***)[***] Different from null of unity at 10%(5%)[1%] percent significance level. Source: Meredith and Chinn (1998).

1 Fixed effects regression. Sample period: 1980Q1-1997Q4.

Table 2. Long-Horizon Tests of Uncovered Interest Parity

$$\Delta s_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^{*}_{t,k}) + \epsilon_{t+k} \,. \qquad \textbf{(7)}$$

Panel 2.a: Benchmark Government Bond Yields, 10-Year Maturity
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$	
German deutschemark	-0.008*** (0.002)	0.829 (0.147)		0.424	
Japanese yen	-0.039*** (0.004)	0.487 (0.101)	***	0.197	
U.K. pound	0.005 (0.004)	0.567 (0.104)	***	0.447	
French franc	-0.007 (0.020)	0.885 (0.508)		0.027	
Italian lira[1]	0.008 (0.007)	0.214 (0.155)	***	0.009	
Canadian dollar	0.000 (0.007)	1.104 (0.657)		0.145	
Constrained panel[2]	...	0.635 (0.111)	***	0.654	...

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample for is 1983Q1-1998Q1. * (***)[***] Different from null of unity at 10%(5%)[1%] percent significance level.

1 Sample period for the Italian lira limited to 1987Q1–1998Q1.

2 Fixed effects regression, excluding the lira. Sample period 1983Q1 - 1997Q4.

Table 2. continued

Panel 2.b: 10-Year Government Bond Yields
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$
Deutschemark	-0.009*** (0.002)	0.836 (0.128)	*	0.468
Japanese yen	-0.038*** (0.005)	0.564 (0.146)	**	0.213
U.K. pound	-0.003 (0.004)	0.719 (0.109)	***	0.446
Constrained panel [1]	...	0.708 (0.090)	***	0.695

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample for is 1983Q1-1998Q1. * (***)[***] Different from null of unity at 10%(5%)[1%] percent significance level.

[1] Fixed effects regression. Sample period: 1983Q1-1997Q4.

Table 2. (continued)

Panel 2.c: 5-Year Government Bond Yields
(MA(19)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$
German deutschemark	0.002 (0.016)	0.914 (0.360)		0.066
U.K. pound	0.002 (0.010)	1.084 (0.415)		0.011
Canadian dollar[1]	0.008 (0.007)	1.337 (0.410)		0.157
Constrained panel[2]	...	1.010 (0.344)		0.396

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample for is 1983Q1-1998Q1. * (***)[***] Different from null of unity at 10%(5%)[1%] percent significance level.

1 Sample period for Canada limited to 1985Q4–1998Q1 (44 degrees of freedom).
2 Fixed effects regression. Sample period 1986Q1 - 1997Q4.

Table 3. Johansen Cointegration Test Results

$$\begin{aligned}\Delta s_t &= \gamma_{10} + \Phi_1[s_{t-1} - \delta_1 f_{t-2} - \delta_0] + \sum_{i=1}^{j} \gamma_{1i}\Delta s_{t-i-1} + \sum_{i=1}^{j} \zeta_{1i}\Delta f_{t-i-2} + \epsilon_{1t} \\ \Delta f_{t-1} &= \gamma_{20} + \Phi_2[s_{t-1} - \delta_1 f_{t-2} - \delta_0] + \sum_{i=1}^{j} \gamma_{2i}\Delta s_{t-i-1} + \sum_{i=1}^{j} \zeta_{2i}\Delta f_{t-i-2} + \epsilon_{2t}\end{aligned} \qquad \textbf{(7)}$$

Panel 3.a: 3 Month Horizon

	LR	δ_0	δ_1	Φ_1	Φ_2	j
Deutschemark	10.10	0.028	1.051 (0.021)	2.232 (0.804) [2.778]	0.005 (0.047) [0.108]	4
Japanese yen	26.76**	0.028	1.006 (0.004)	3.477 (1.025) [3.392]	0.370 (0.095) [3.901]	2
U.K. pound	19.36**	-0.006	1.015 (0.012)	2.772 (1.030) [2.692]	0.177 (0.073) [2.438]	3
French franc	14.84	-0.266	0.851 (0.180)	-0.798 (0.260) [3.069]	-0.018 (0.028) [0.650]	4
Italian lira	18.95**	0.362	1.050 (0.038)	1.355 (0.665) [2.038]	0.322 (0.131) [2.451]	10
Canadian dollar	10.43	0.013	1.056 (0.026)	1.003 (0.544) [1.843]	0.154 (0.070) [2.196]	3

Notes: LR is the likelihood ratio for the Maximal Eigenvalue test of the H0 of zero cointegrating vectors against HA of one cointegrating vector. 15.41 and 20.04 are the 5% and 10% critical values (Osterwald and Lenum, 1992). Point estimates from Johansen maximum likelihood procedure. *j* is the number of lags in the VAR representation of the cointegrated system. Sample for is 1980Q1-1998Q1. * (***)[***] Different from null of unity at 10%(5%)[1%] percent significance level.

Panel 3.b: 5 Year Horizon

	LR	δ_0	δ_1	Φ_1	Φ_2	j
Deutschemark [1]	12.73	0.699	2.008 (1.274)	-0.037 (0.019) [1.939]	0.045 (0.020) [2.243]	8
U.K. pound [2]	14.15	0.281	0.477 (0.233)	-0.173 (0.067) [2.573]	0.025 (0.146) [0.172]	13
Canadian dollar [3]	--	--	1	-0.058 (0.028) [2.077]	-0.075 (0.047) [1.585]	5

Notes: LR is the likelihood ratio for the Maximal Eigenvalue test of the H0 of zero cointegrating vectors against HA of one cointegrating vector. 15.41 and 20.04 are the 5% and 10% critical values (Osterwald and Lenum, 1992). Point estimates from Johansen maximum likelihood procedure, standard errors in parentheses, absolute values of the *t*-statistics in brackets. *j* is the number of lags in the VAR representation of the cointegrated system.
* (***)[***] Different from null of unity at 10%(5%)[1%] percent significance level.

[1] Sample is 1980Q1-98Q1.
[2] Sample is 1978Q1-98Q1.
[3] Sample is 1987Q1-98Q1. Long run unbiasedness imposed in Horvath-Watson regression.

Table 4. Simulation Model

Uncovered interest parity:

$$\Delta s^{e}_{t,t+1} = \hat{i}_t - \eta_t$$

Monetary reaction function:

$$\hat{i}_t - \hat{\pi}_t = 0.5\,(\hat{\pi}_t + \hat{y}_t)$$

Inflation (π) equation:

$$\hat{\pi}_t = 0.6\,\hat{\pi}_{t-1} + (1-0.6)\,\hat{\pi}^{e}_{t,t+1} + 0.25\hat{y}_t + 0.1\,\Delta(s_t - \hat{p}_t) + \nu_t$$

Output (y) equation:

$$\hat{y}_t = 0.1\,(s_t - \hat{p}_t) - 0.5\,(\hat{i}^{l,e}_{t} - \hat{\pi}^{l,e}_{t}) + 0.5\,\hat{y}_{t-1} + \epsilon_t$$

Price level (p) identity:

$$\hat{p}_t = \hat{p}_{t-1} + \hat{\pi}_t$$

Exchange rate (s) identity:

$$s_t = s_{t-1} + \Delta s_t$$

Long-term expected interest rate:

$$\hat{i}^{l,e}_{t} = (1/5)\,(\hat{i}_t + \hat{i}^{e}_{t,t+1} + \hat{i}^{e}_{t,t+2} + \hat{i}^{e}_{t,t+3} + \hat{i}^{e}_{t,t+4})$$

Long-term expected inflation rate:

$$\hat{\pi}^{l,e}_{t} = (1/5)\,(\hat{\pi}_t + \hat{\pi}^{e}_{t,t+1} + \hat{\pi}^{e}_{t,t+3} + \hat{\pi}^{e}_{t,t+3} + \hat{\pi}^{e}_{t,t+4})$$

Note: ^ denotes a variable expressed relative to the US.

Table 5. Regression Results from Stochastic Simulations

	Regression horizon: 1 year	5 years	10 years
Estimated slope coefficient (β)			
Average	-0.50	0.82	0.78
Maximum	0.23	1.32	1.21
Minimum	-1.37	0.53	0.30
Standard deviation	0.37	0.17	0.21
Standard error of β			
Average	0.42	0.18	0.18
Maximum	0.53	0.22	0.22
Minimum	0.31	0.14	0.14
Standard deviation	0.04	0.02	0.02
Adjusted R^2			
Average	0.01	0.21	0.19
Maximum	0.08	0.46	0.44
Minimum	-0.01	0.06	0.03
Standard deviation	0.03	0.09	0.10

Figure 1. Recursive Coefficients for 3 Month Horizon

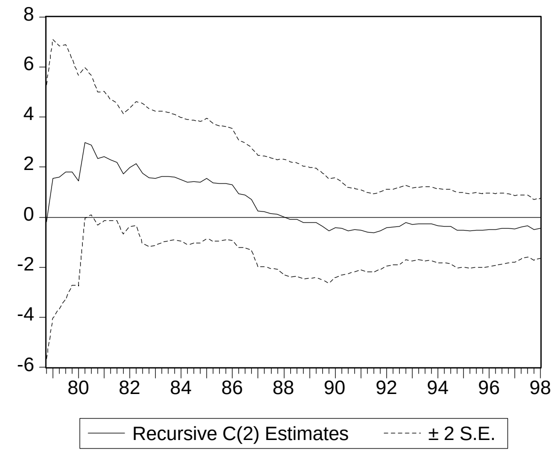

Figure 1.a: Canadian dollar

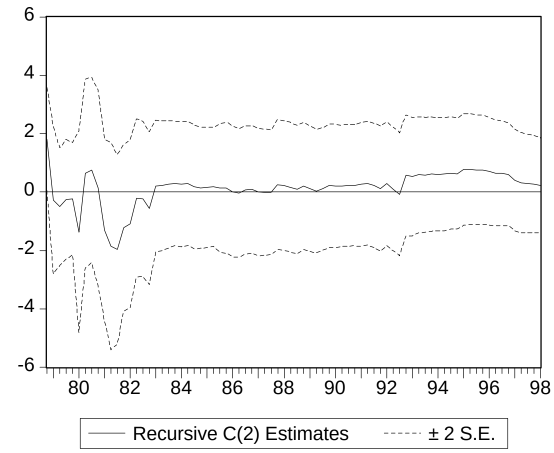

Figure 1.b: French franc

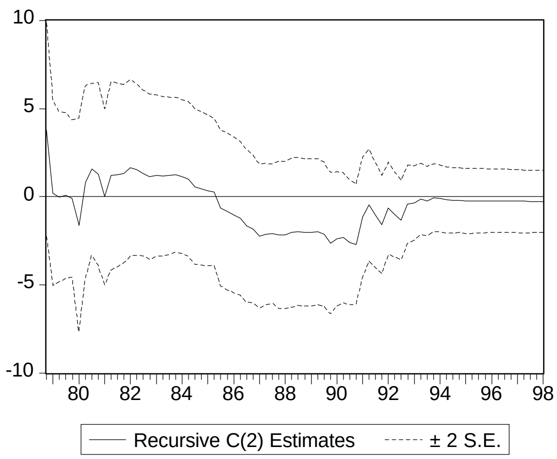

Figure 1.c: Deutschemark

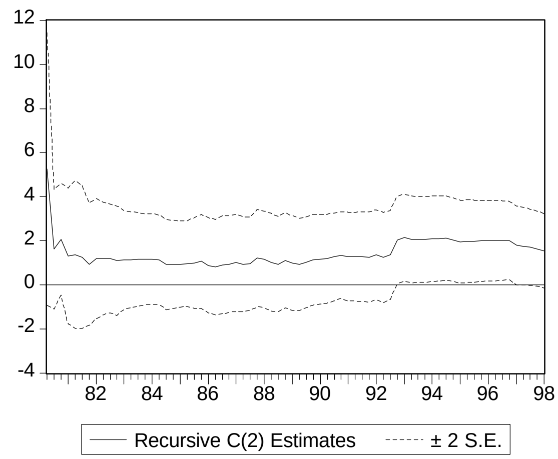

Figure 1.d: Italian lira

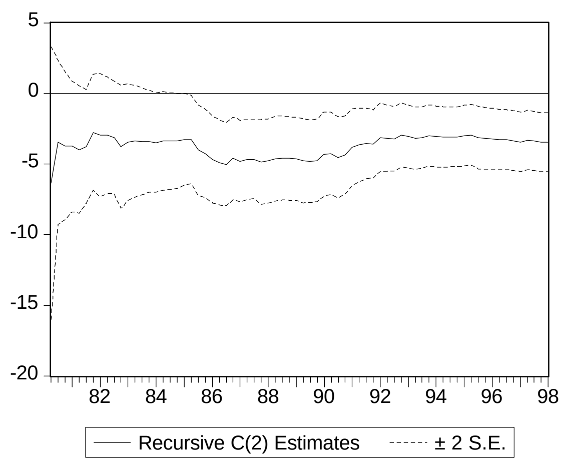

Figure 1.e: Japanese yen

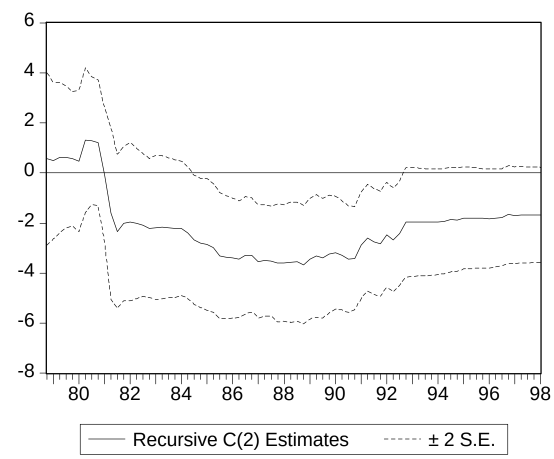

Figure 1.f: U.K. pound

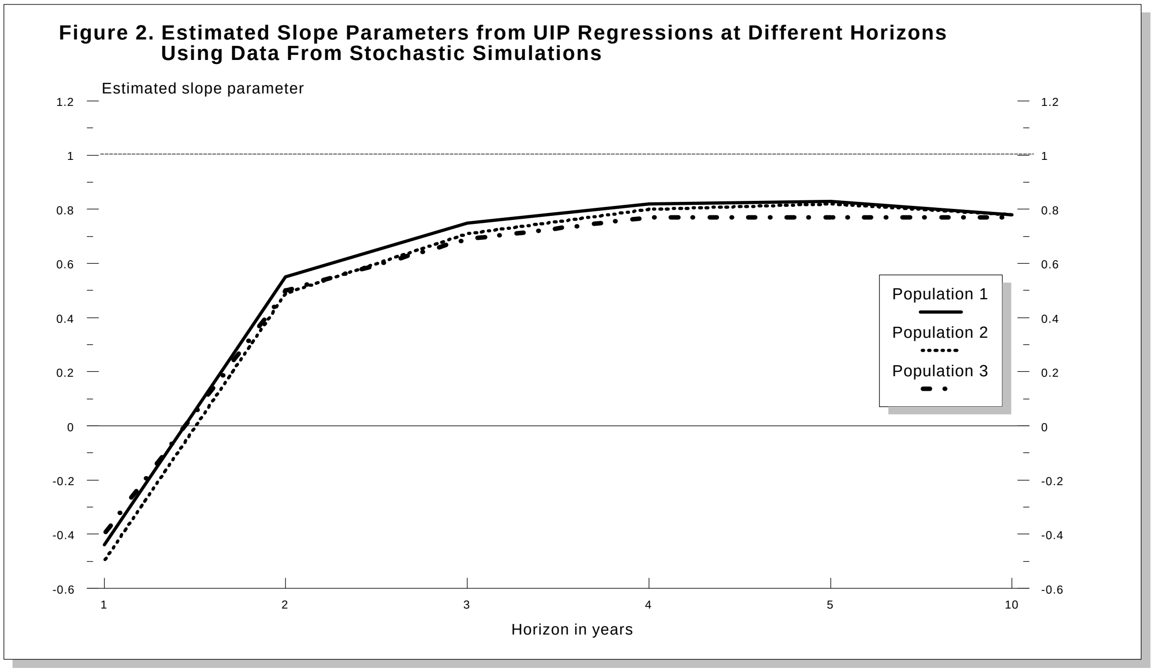
Figure 2. Estimated Slope Parameters from UIP Regressions at Different Horizons
Using Data From Stochastic Simulations
Estimated slope parameter
1.2
1
0.8
0.6
0.4
0.2
0
-0.2
-0.4
-0.6
Population 1
Population 2
Population 3
1
2
3
4
5
10
Horizon in years